

Time Dilation of BATSE Gamma-ray Bursts

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ABSTRACT

We perform several nonparametric correlation tests on the BATSE 3B data to search for evidence of cosmological time dilation. These tests account for the effects of data truncation due to threshold effects in both limiting brightness and limiting duration, enabling us to utilize a larger number of bursts than in previous analyses. We find little significant evidence for correlation between various measures of peak intensity and duration, but the tests cannot conclusively rule out time dilation factors of 2 or less without more data. There is stronger evidence for a positive correlation between fluence and duration, which if confirmed would rule out simple no-evolution cosmological models unless there is a strong intrinsic correlation between the total radiant energy and the duration of bursts.

Subject headings: gamma rays: bursts

1. Introduction

Observations taken by the BATSE instrument aboard the *Compton Gamma-Ray Observatory* have shown that the angular distribution of gamma-ray bursts (GRBs) is isotropic (Meegan et al. 1992), while the $\log N$ - $\log S$ distribution flattens for weaker bursts. Such a distribution would be expected if bursts were of cosmological origin. Paczyński (1992) and Piran (1992) suggested that if bursts originated at cosmological distances, their light curves should be stretched due to cosmological time dilation. If bursts were standard candles, dimmer (hence more distant) bursts would be time-dilated more than brighter (hence less distant) bursts, by a dilation factor $(1 + z_{dim})/(1 + z_{bright})$, where z_{dim} and z_{bright} are the redshifts.

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However, the expected redshift range of order unity would result in a time dilation factor of a few while the burst durations cover a large dynamic range from tens of milliseconds to hundreds of seconds (Kouveliotou et al. 1993). Therefore a time dilation effect can only be detected statistically. Norris et al. (1994, 1995) (hereafter N1 and N2) searched for time dilation effects by dividing the bursts into groups based on their peak count rate C_P and comparing some measure of burst duration with peak count rate. They found evidence that brighter bursts had shorter durations than dimmer ones, and that the difference between the average durations of bright and dim bursts was consistent with a time dilation factor of about 2. This factor of 2 time dilation is what would be expected for simple cosmological models if the peak luminosity of the bursts were independent of their duration or redshift. This effect would be easier to detect if the peak luminosity function of the bursts was narrow; in other words, if the burst peak luminosity was a “standard candle.” Fits of the $\log N$ - $\log C_P$ distributions to cosmological models with the standard candle peak luminosity assumption (Piran 1992; Mao & Paczyński 1992; Wickramasinghe et al. 1993; Fenimore et al. 1995; Azzam & Petrosian 1996) obtain a mean redshift $z_{dim} \sim 1$ so that the dilation factor would be about 2 ($z_{bright} \ll 1$).

However, Band (1994) has cautioned that an intrinsic burst luminosity function could easily produce similar effects. Yi & Mao (1994) also noted that relativistic beaming in either Galactic halo or cosmological models can produce flux-duration relationships that might be consistent with the reported effects. Wijers & Paczyński (1994) suggested a way to distinguish between anticorrelations between flux and duration produced by cosmological time dilation and those produced by a decrease in burst density with distance, which is needed in a local extended halo model if the shape of the luminosity function is independent of distance. They cautiously concluded that the data from the first BATSE catalog is more consistent with a cosmological interpretation. On the other hand, as shown by Fenimore (1996), the agreement of the time dilation results with the cosmological models is destroyed by the fact that the burst duration T (or the pulse widths) appears to be well correlated with photon energy E ; $T \propto E^{-\alpha}$ (Fenimore et al. 1995). Since higher redshifts correspond to higher energies of emitted photons, the expected time dilation effect will not yield a $T \propto (1+z)$ relation but rather $T \propto (1+z)^{1-\alpha}$. For $\alpha = 1/2$ this would require higher redshifts ($z_{bright} \approx 1$ and $z_{dim} \approx 6$), destroying the good agreement with the $\log N$ - $\log S$ results. Furthermore, there is even now disagreement about the reality of the observed time dilation. For example, Mitrofanov et al. (1994; 1996) compared averaged time histories of weak and strong bursts detected by BATSE and found no evidence for any time dilation. Similarly, Fenimore et al. (1995) noted that a different method of analysis, which agrees with the results of N1 and N2 when used for the same set of data, gives a much smaller time dilation factor (1.3 instead of 2) for a larger set of data. Recently, Norris et al. (1996) has

reported that the larger data set is consistent with observed time dilation factors of 1.5–1.7.

It is clear that despite the numerous works published on the subject, time dilation of gamma-ray bursts remains controversial. Here we present new results on this topic which differ from previous works in two important ways. Previous studies of burst time dilation have been limited to bright, long duration bursts. This selection avoids the biases against detection by BATSE of bursts with durations shorter than the trigger integration time, and against weaker bursts due to the variability of the threshold photon count rate. As a result, these studies utilize only a small percentage ($\sim 20\%$) of the total number of bursts. In this work we attempt to extend the time dilation analyses to a sample of bursts which is larger by about a factor of two, by properly accounting for the effects of variable thresholds and the short duration bias. These biases and the methods to account for them are described in detail in several of our previous publications (Efron & Petrosian 1992; Petrosian 1993; Lee, Petrosian, & McTiernan 1993; Petrosian, Lee, & Azzam 1994). For a complete review see Lee & Petrosian (1996; hereafter LP).

Second, all of the previous studies use the peak photon count rate or the peak flux as a measure of the distance. This usage assumes that the peak photon luminosity is independent of the burst duration or distance. In addition, a narrow distribution of the peak luminosity is required for the detection of the small time dilation effect. In other words the peak photon luminosity should be nearly a standard candle for the previous studies to be valid. This seems not very likely considering the large dispersion in the duration and pulse shapes of GRBs. We believe that it is more likely that the total energy (or total number of photons) emitted by a burst is a standard candle, so that the energy or photon fluence will be a better measure of the distance to the bursts. Because bursts are generally detected on the basis of their peak flux, it has been difficult in the past to examine unbiased fluence distributions. However, as shown by Petrosian & Lee (1996; hereafter PL) the methods mentioned above work equally well with fluence, so we carry out our tests using both peak fluxes and fluences. Furthermore, because of the complex and varied burst pulse shapes, it is not clear what measure of the time structure, or which of the several available time scales associated with the pulse profiles, would be a reliable measure of redshift in a cosmological scenario. In addition to using the durations T_{50} and T_{90} provided in the BATSE catalog, we also use an effective duration defined by the ratio of the total energy released to the peak luminosity, which we believe to be a robust measure of duration and therefore redshift. A brief summary of our method is given in §2, the choice of test variables is discussed in §3, the test results are presented in §4, and a discussion and summary of our conclusions are given in §5.

2. Analysis Method

The problem of searching for time dilation effects can be thought of as a search for a correlation between two variables, one of which is some measure of burst duration while the other is some measure of burst brightness. Due to the detection biases against short duration and weak bursts, the practice in previous dilation studies has been to examine relatively bright, long duration bursts.

However, it is possible to extend such a test to a much larger sample if the observational selection criteria or data truncations are well defined, because there exist methods to test for correlations in the presence of such truncations. A simple test which is easily applied to burst data is the t_w test described by Efron & Petrosian (1992). Briefly, the test relies on the concept of the *associated set* of points for each data point. For an untruncated data set, the associated set of data points is the entire set of data points. For a truncated data set, the associated set differs for each data point and is defined as the largest subset of data points for which there is no truncation. Each data point can then be assigned a rank amongst the points comprising its associated set. If a correlation exists, then the ranks will be correlated. Because the expected distribution of ranks for an uncorrelated data set is well-defined no matter what the individual distribution functions of the variables are, the test is completely nonparametric. Efron & Petrosian’s t_w test has the following properties:

1. The scalar statistic t_w calculated from the data could be positive or negative, with the sign indicating correlation (+) or anticorrelation (-).
2. The value of t_w gives the probability $P(t_w) = \text{erfc}(|t_w|/\sqrt{2})$ that the data were drawn from an uncorrelated distribution. In other words, t_w is distributed normally and can be interpreted as the number of standard deviations away from the expected result of the test for perfectly uncorrelated data. In this sense t_w can be used as an error estimate.
3. Each data point is weighted, and the weights can be chosen such that the t_w test becomes equivalent to standard tests such as Kendall’s τ for simple truncations. Throughout this paper we give every data point the same weight, which we indicate notationally by referring to this variation of the t_w test as t_1 .

Thus, to test the time dilation hypothesis, it becomes merely a matter of choosing the variables to test and defining the data truncations. As discussed below, these tasks can be far from straightforward.

3. Choice of Variables

The BATSE catalog provides several observed quantities which could be used as a measure of distance to the burst. It is not immediately obvious which measures are the most suitable, so we consider them carefully below.

3.1. Peak Photon Flux

The most observationally tractable quantities are the average peak photon count rate $\bar{C}_P = C_{max}/\Delta t$, where the average is over the trigger time Δt . The burst selection criterion is that $\bar{C}_P > \bar{C}_{lim}$, where $\bar{C}_{lim} = C_{min}/\Delta t$ is the threshold value set by the background count rate. The BATSE catalog gives the values of C_{max} and C_{min} for three trigger times $\Delta t = 64, 256$, and 1024 ms. The rate \bar{C}_P is related to the peak photon luminosity through the instrument response and its dependence on the angular location of the burst. A better measure is the average peak photon flux \bar{f}_P , which is directly related to the average peak photon luminosity \bar{F}_P as $\bar{f}_P = \bar{F}_P/4\pi d_L^2$, where d_L is the appropriate luminosity distance. In the cosmological scenarios d_L depends on the redshift and the model parameters. Thus, in the rest frame of any particular burst the time interval over which the peak flux is averaged depends on its redshift. As described in LP the threshold flux for each burst can be obtained from the catalog as

$$\bar{f}_{lim} = \bar{f}_P(\bar{C}_{min}/\bar{C}_P) = \bar{f}_P(C_{min}/C_{max}). \quad (1)$$

These time-averaged flux measures are subject to the short duration bias for bursts with durations less than the trigger interval Δt . It was shown in LP (see also Petrosian, Lee, & Azzam 1994) that an approximate way of correcting for this bias is to define our best estimate of the “true” or instantaneous peak flux as

$$\hat{f}_P = \bar{f}_P(1 + \Delta t/T), \quad (2)$$

where T is the “true” duration of the burst. Typically some measure of the duration is used to estimate this true duration, so that equation (2) becomes

$$\hat{f}_P(T_x) = \bar{f}_P(1 + \eta\Delta t/T_x), \quad (3)$$

where T_x is some observationally tractable duration measure and $\eta \equiv T_x/T$. LP found that $\eta \approx 0.5$ fit the BATSE data for $T_x = T_{50}$ in a statistical sense. A similar correction can be used to define an estimate of the instantaneous peak photon count rate \hat{C}_P from the observed average value \bar{C}_P . Note that these transformations remove not only the duration bias but also the ambiguity of the redshift dependence of the rest frame trigger interval. In the equation $\hat{f}_P = \bar{F}_P/4\pi d_L^2$, the only redshift dependence is through d_L .

3.2. Fluence

As mentioned before it seems unlikely that F_P is a standard candle. Another possible candidate is the total radiated energy ε or the total number of emitted photons. Unlike F_P , neither of these quantities would be affected by the boosting due to the large bulk Lorentz factors that would be appropriate for a cosmological fireball scenario (Paczynski 1986; Goodman 1986; Meszaros & Rees 1993, 1994). The appropriate observational measure found in the BATSE catalog is the energy fluence $\mathcal{F} = \varepsilon/4\pi\tilde{d}_L^2$, where $\tilde{d}_L = d_L/\sqrt{1+z} = d_m(1+z)$ where d_m is the metric distance. As described in PL the threshold on the fluence is obtained as

$$\mathcal{F}_{lim} = \mathcal{F}(C_{min}/C_{max}). \quad (4)$$

Like \hat{f}_P and C_P , the fluence \mathcal{F} is also subject to a bias. The bias now is against the detection of weak and long bursts which have \bar{C}_P too low to exceed the threshold \bar{C}_{lim} . Following the same arguments used in LP which led to the correction of peak fluxes according to equation (2), it can be shown that

$$\mathcal{F}_{lim} \approx \langle h\nu \rangle \bar{f}_{lim}(\Delta t + T), \quad (5)$$

where $\langle h\nu \rangle$ is the average energy per photon in the 50–300 keV range. Because of the strong dependence of \mathcal{F}_{lim} on T , we argue later in §4.2.2 that a test of correlation between \mathcal{F} and \mathcal{F}_{lim} can be used to effectively test the duration–fluence correlation.

3.3. Observational Measures of Duration

3.3.1. Effective Duration

As a measure of redshift we would ideally like to use the “true” duration T of the burst, which is an observationally ill-defined quantity. There are several available observational measures T_x of burst duration which could potentially be used as measures of redshift. Two of these, T_{90} and T_{50} , are listed in the catalog and represent the time interval between the instances when the burst reaches 5% and 95% of its total fluence for T_{90} , and 25% to 75% of the fluence for T_{50} . As an alternative measure of duration we define an “effective duration”

$$T_{eff}(T_x) = \frac{\mathcal{F}}{\hat{f}_P \langle h\nu \rangle} = \frac{\mathcal{F}}{\bar{f}_P (1 + \eta \Delta t / T_x) \langle h\nu \rangle}. \quad (6)$$

We calculate $\langle h\nu \rangle$ by assuming a power law energy spectrum and using the ratio of 100–300 keV to 50–100 keV fluence to solve for the power law index. This approximation should not affect the correlation analysis described below unless there is a very strong correlation between the spectral index and fluence or flux. In any case, the value of $\langle h\nu \rangle$ is insensitive

to the spectrum, varying by less than 10% for simple power laws and by less than 25% for Band (1993) type spectra. We note, however, that there does exist a correlation between spectral hardness and duration, in that shorter bursts tend to have harder spectra (Kouveliotou et al. 1993). If bursts were cosmological, their spectra should be more redshifted the more distant they are. Since a “typical” GRB tends to have a spectral steepening at about 150 keV (Band et al. 1993), a redshifted burst will be expected to have a softer spectrum, and hence a smaller $\langle h\nu \rangle$ than a local burst. Not correcting for this effect can only accentuate the time dilation effect on the effective duration, because the smaller values of $\langle h\nu \rangle$ for bursts at high redshift will tend to make T_{eff} longer due to the form of equation (6). As will be seen below, we do not believe that the results of the correlation tests warrant a correction for the effect.

3.3.2. Signal-to-Noise Bias

Norris (1996) has shown that there are problems with using T_{50} and T_{90} as estimators of the time dilation factors. Consider bursts with light curves consisting of a dominant spike and one or more smaller spikes separated by long quiescent periods (for example, see the light curves of bursts 143, 219, 841, 1145, 1440 in the BATSE catalog). Such bursts will be assigned long values of T_{90} due to the presence of the small spike(s). However, for bursts of similar time profiles but weaker intensities these small spike features could be lost in the background noise and would be considered as single spike bursts, leading to much shorter values of T_{90} . This signal-to-noise bias would lead to a *correlation* between C_P and T_{90} , and would tend to cancel out a time dilation effect. To a lesser extent, T_{50} suffers from this bias as well. We believe T_{eff} may be a more stable measure of duration with respect to fluctuations in the background noise level. For instance, for multi-peaked bursts such as those mentioned above both T_{50} and T_{90} could jump discontinuously as a function of signal-to-noise ratio, while T_{eff} would change gradually and often insignificantly. On the other hand, T_{eff} does suffer from the necessity of estimating the peak flux from time-averaged data, requiring some independent measure of duration (e.g. using Eq. [2] with $2T_{50}$ in place of T). However, this is a second-order effect because for the majority of bursts $T_x \gg \eta\Delta t$ and hence T_{eff} is independent of T_x .

Figure 1 shows a comparison of the various definitions of duration obtained with BATSE catalog data. In these plots, we show ratios of the various measures of duration as functions of \bar{C}_P/\bar{C}_{lim} , which should serve as an estimate of the signal-to-noise ratio. In the top two panels of the Figure, it can be seen from the graphs that there is a systematic trend for bursts with smaller signal-to-noise ratio to have smaller ratios of T_{90}/T_{eff} and T_{90}/T_{50} ,

clearly demonstrating the signal-to-noise bias. This effect can approach a factor of 2, which is of the same order as the expected time dilation effect. As shown in the bottom panel, T_{50} is much less susceptible to the bias than T_{90} . We have repeated this analysis for a subset of bursts kindly provided to us by J. Norris, for which the T_{50} and T_{90} durations have been calculated only after normalizing the peak intensities in order to remove the signal-to-noise bias. We denote these corrected duration estimates as \hat{T}_{50} and \hat{T}_{90} . Substituting these peak-normalized durations for T_{eff} in the ratios depicted in the Figure, we find trends similar to those shown in Figure 1. Comparing T_{eff} to the peak-normalized measures of duration seems to show very little evidence for this trend, in agreement with our conclusion above that T_{eff} should not suffer as much from the bias.

Clearly, it would be preferable to use peak-normalized burst duration estimates \hat{T}_{90} and \hat{T}_{50} instead of the BATSE values of T_{90} and T_{50} in our t_w test. However, the very utility of our test lies in its ability to extend the burst sample to include very weak bursts. It is precisely these bursts for which the peak-normalization procedure becomes problematic. Lacking \hat{T}_{90} and \hat{T}_{50} for the larger sample of bursts we wish to test, we take T_{eff} to be the most robust available duration measure. We also note that in LP it was shown that corrections for the short duration bias based on equation (3) provided the best agreement between $\Delta t = 64$ ms and $\Delta t = 1024$ ms duration distributions when T_{eff} was used as an estimator for T_x .

4. Correlation Test Results

We have performed the t_w test for stochastic independence on several combinations of variables. In all these tests, we choose the standard weight vector $\vec{w} = (1, 1, \dots, 1)$, so that each data point is given the same relative importance (see Efron & Petrosian 1992 for a discussion of weights in the t_w test). As a measure of distance we use the average peak flux \bar{f}_P , the estimated true peak flux \hat{f}_P , and the fluence \mathcal{F} . We do not consider \bar{C}_P and \hat{C}_P , which give results similar to \bar{f}_P and \hat{f}_P , respectively. As a measure of redshift we use the durations T_{90} , T_{50} , and $T_{eff}(T_{50})$ as defined above. We carry out these tests for data obtained at all three trigger times $\Delta t = 64, 256$, and 1024 ms separately.

4.1. Direct Test for Peak Flux vs. Duration

A direct test of the correlation between the average peak flux \bar{f}_P and duration is complicated because the threshold \bar{f}_{lim} varies and a clear truncation boundary cannot be

delineated in the \bar{f}_P – T plane. Analysis of the three-dimensional distribution involving \bar{f}_P , \bar{f}_{lim} , and T is required to properly take into account the variation of \bar{f}_{lim} . This point is discussed in §4.2, but for the sake of simplicity and clarity we first limit ourselves to the two-dimensional case by selecting a subsample of data which could be described with a single constant threshold \bar{f}_{lim} . An obvious choice is to limit the subsample to the sources with \bar{f}_P greater than or equal to the maximum value of the observed values of \bar{f}_{lim} . In this case, all sources with $\bar{f}_P \leq \bar{f}_{lim,max}$ are excluded from the analysis. A slightly better choice is to find the value of $\bar{f}_{lim} = \bar{f}_{lim,0}$ such that the truncation is kept simple and the number of data points is maximized. In practice this amounts to limiting the data points to those with $\bar{f}_{lim} \leq \bar{f}_{lim,0}$ and $\bar{f}_P \geq \bar{f}_{lim,0}$. With a constant \bar{f}_{lim} , the truncation boundaries in the \bar{f}_P – T plane become parallel to the axes and the problem reduces to one of simple truncation. The t_1 test then reduces down to a simple rank order correlation test.

4.1.1. Average Peak Flux—Duration Correlation

Using this truncation, we calculate $t_1(\text{data})$ values to test the correlation between the average peak flux \bar{f}_P and various observational measures of duration. The first three rows of Table 1 show the values of $t_1(\text{data})$ for each of the three trigger times $\Delta t = 64, 256$, and 1024 ms, respectively. The first obvious feature in these numbers is that the values of $t_1(\text{data})$ are significantly and consistently larger for correlations involving T_{90} than for the other measures of duration, which give nearly identical results. The most likely explanation of this result is that the T_{90} values are underestimated at low values of \bar{f}_P due to the signal-to-noise bias, giving rise to a larger positive value for $t_1(\text{data})$ and an apparent correlation. This result is in agreement with the relations discussed in connection with Figure 1 and the findings by Norris (1996).

4.1.2. Corrected Peak Flux—Duration Correlation

The second feature of these $t_1(\text{data})$ values is that they are larger for larger values of Δt . This result is most likely due to the short duration bias mentioned in connection with equation (2). We may use the approximation of equation (2) to correct the average peak flux, using $(100/x)T_x$ as estimates for the true duration T . The magnitude of this correction increases with the ratio of $\Delta t/T$, and is therefore largest for $\Delta t = 1024$ ms at $T \lesssim 1$ s. The transformed data in the \hat{f}_P – T_x plane is no longer truncated by a single average peak flux limit. Instead the truncation is defined by

$$\hat{f}_P > \hat{f}_{lim,0} = \bar{f}_{lim,0}(1 + \Delta t/T), \quad (7)$$

which cannot be described as a simple truncation (see Fig. 6 of LP). With proper account for the truncation given by equation (7), the values of $t_1(\text{data})$ obtained for these data are given in the second three rows of Table 1. The values of $t_1(\text{data})$ are now lower, especially for the $\Delta t=1024$ ms data. This result apparently confirms the assertion that the differences between the data sets corresponding to the three different values of Δt are due to the short duration bias. In particular the good agreement found for T_{eff} and the three values of Δt indicates again that equations (2), (3), and (7) are fairly reliable. The least biased of these tests should be the one involving $T_{eff}(T_{50})$ and $\hat{f}_P(T_{eff})$ with $\Delta t = 64$ ms. This test resulted in a value of $t_1 = -0.565$, which corresponds to rejection of the null hypothesis of independence at a 43% probability. Therefore, the test is consistent with no correlation, although it may also be consistent with the weak anticorrelation expected from time dilation (see §4.1.5). The results of all of the tests are summarized graphically in Figure 2.

4.1.3. Importance of the Signal-to-Noise Bias

A correction for the signal-to-noise bias could change these numbers. Table 2 shows the results of similar tests performed only on the subset of bursts for which Dr. J. Norris kindly supplied us with peak-normalized durations. This subset consisted of 265 bursts with long ($T \gtrsim 1$ s) durations and peak count rates greater than 1400 counts s^{-1} , so that the short duration bias and the problems associated with the variable threshold rate are minimized. As expected, tests involving the BATSE T_{90} give much more positive correlation test results than those involving the peak-normalized duration \hat{T}_{90} . The same trend, but at a much less statistically significant level, can be seen in the \hat{T}_{50} results. It is interesting that this subset of bursts gives significantly more negative results than the tests involving the larger sample of bursts shown in Table 1, no matter which measure of duration is used. Since the peak-normalized subsample contains mostly long duration events, this result could be an indication that long and short duration bursts have different correlation trends. In any case, the value of $t_1(\text{data}) = -1.61$ for the test involving $T_{eff}(T_{50})$ and $\hat{f}_P(T_{eff})$ for this sample of long duration bursts is on the verge of being significant, with the probability of rejecting the null hypothesis at 89%. All other numbers indicate less significant rejections, implying weaker or no anticorrelations.

4.1.4. Test for Correlation Trends

It is possible that the less significant results for the larger sample could arise from correlations of opposite trends in different portions of the data. For example equal and

opposite relations between \hat{f}_P and say short ($T < 1$ s) and long ($T > 1$ s) bursts can cancel each other out, giving low $t_1(\text{data})$ values. We test this possibility by dividing the data into subsamples.

Figure 3 shows the values of $t_1(\text{data})$ as a function of cutoff in $T_{eff}(T_{50})$. Both maximum and minimum cutoffs for $T_{eff}(T_{50})$ are shown, so that for any subdivision of the data into duration-limited subsets one may read off the $t_1(\text{subset})$ values for those subsets. Thus for a given value of T_{eff} the middle (lower) panels give the value for $t_1(\text{subset})$ for all T_{eff} less (greater) than the specified value. The dotted line shows $-0.565\sqrt{m/296}$, where m is the number of points in the subset. Because of the way t_1 is defined (see Efron & Petrosian 1992 for details), if the average normalized rank of each point is the same an increase (or decrease) in the number of points should lead to an increase (or decrease) in the magnitude of t_1 by about a factor of $\sqrt{m/M}$, where M is the original number of points. There are two areas where the data seem to deviate from this form. For bursts above $T_{eff}(T_{50}) \approx 10$ s the $t_1(\text{subset})$ values appear to reach significantly negative values ($t_1 < -1.645$), while for bursts below $T_{eff}(T_{50}) \approx 0.1$ s the $t_1(\text{subset})$ values approach significantly positive values ($t_1 > 1.645$). This finding is consistent with the difference in t_1 values given in Tables 1 and 2. Such behavior could be an indication that short and long duration bursts have different correlation properties, which might be the case if long duration bursts showing the time dilation were cosmological and short duration bursts consistent with no time dilation (or even possibly time contraction) were local. However, the number of bursts in each of these subsets is small, so that chance cannot be ruled out as a reason for these differences.

4.1.5. Interpretation of the Test Results

The overall test results may be consistent with no correlation, but they may also be consistent with a very weak anticorrelation. Since the expected time dilation factors are of order 2 or less while the range of durations spans many orders of magnitude, the question becomes “what value of t_1 would we expect given such a weak correlation?” To answer this question we rely on simulations. We create a large number of simulated durations and peak fluxes, chosen such that the univariate distributions of T and \hat{f}_P closely resemble those actually observed (and shown in LP). For our reference simulations we chose \hat{f}_P such that its differential distribution followed a broken power law with a logarithmic slope of -2.5 above a flux of 20 photons $\text{cm}^{-2} \text{s}^{-1}$. At lower fluxes, the logarithmic slope flattens to -2.0. Independently, T was chosen to be distributed as two lognormal peaks at 0.1 and 10 seconds, with a standard deviation of one order of magnitude. The selection effects which

were appropriate for the real data were applied to the simulated data, and the number of data points in each simulation was chosen to match the number of untruncated real data points (296). The statistic $t_1(\text{simdata})$ was calculated for each realization and a distribution of t_1 was formed. Figure 4 shows this distribution as the solid histogram. As expected, the distribution is approximately normal with mean 0 and variance 1. We then repeated the simulations, adding a small amount of anticorrelation in a power-law fashion such that the expected observed time dilation factors for the brightness bins analogous to those chosen by Norris et al. (1994) were 1.3, 1.7, and 2.0. The distributions of t_1 for these simulations are shown as the long-dashed, dotted, and short-dashed histograms, respectively (their ordinates have been shifted for clarity). The mean values of the distributions are -0.127, -0.67, -1.21, and -1.51. Clearly, there is significant overlap of all of the distributions, so that a test result of t_1 between about 0 and -1.6 would be consistent with any of these at about the 10% level of confidence. The conclusion therefore must be that with the current number of data points, the test cannot at present distinguish conclusively between no correlation and the weak correlation signature expected from cosmological time dilation. Assuming a very simple correlation form such that the average normalized rank of each point is the same, in order for a $t_1 = 0$ result to rule out an observed factor of 2 dilation to the 90% confidence level, one would need $296(1.645/1.51)^2 \approx 350$ bursts in the sample. For a $t_1 = 0$ result to rule out an observed factor of 1.3 dilation to the same level of confidence requires about 1800 bursts!

4.1.6. Error Correlation Bias

There may exist yet another bias in all of these tests, which arises because the errors on quantities such as \hat{f}_P may be correlated with the values of T due to the dependence of the \hat{f}_P on T . For example, such a correlation might increase the spread in \hat{f}_P with decreasing T , which could lead to an edge effect due to the non-symmetric shape of the \hat{f}_P distribution. We have attempted to estimate the magnitude of this effect on the measured t_1 values through simulations. We create a number of simulated \hat{f}_P - T data sets using the method described in §4.1.5 and calculate the values of $t_1(\text{simdata})$. We then randomly shift the data points by an amount which depends on their particular values of \hat{f}_P and T which is determined by the observational errors in the actual BATSE data. Due to this shifting, some points move across the data truncation boundary into the observable set of data points, while others become unobservable. Note that by using this procedure we also take into account the “peak flux bias” caused by Poisson fluctuations in weak bursts. $t_1(\text{simdata}')$ is recalculated for this error-shifted data set, and the difference between it and the original statistic is found. The process is repeated for a large number of simulated data

sets. Fortunately, for the specific error properties and data truncations relevant for the BATSE data it turns out that the largest difference between the average t_1 for the reference simulations and the average t_1 for the error-shifted simulations is small (< 0.1). This effect tends to give error-shifted t_1 values that are slightly larger than the true t_1 values. Note that any correction for this effect and for the correlation between duration and spectral index mentioned previously in §3.3.1 would tend to reduce the significance of any time dilation signature. Given that the significance of the test results is marginal to begin with, we do not consider correcting the results for these additional effects.

4.2. Indirect Tests of Correlation

It is possible to use the complete three dimensional data set \hat{f}_P , \hat{f}_{lim} , and T_x to determine the correlation between any of the three variables. The threshold flux \bar{f}_{lim} set prior to the occurrence of a burst is expected to be uncorrelated with any of the burst properties, in particular \bar{f}_P , \hat{f}_P , or T_x . As we shall see below this does not seem to be true for all cases. For the moment assuming the expected absence of correlation, an analysis of the three dimensional data effectively gives the correlation between \hat{f}_P and T_x . The description of such an analysis is complicated but it turns out to be unnecessary because it can easily be reduced to the simpler two dimensional case. This is accomplished by the transformation of \bar{f}_{lim} and T_x into \hat{f}_{lim} as described by equation (7). With this transformation we now have a two dimensional distribution of \hat{f}_P and \hat{f}_{lim} with the simple truncation $\hat{f}_P \geq \hat{f}_{lim}$. Therefore, a test of the correlation between \hat{f}_P and \hat{f}_{lim} can be made without the exclusion of some of the data that was necessary to produce Tables 1 and 2. If the \hat{f}_P – \hat{f}_{lim} test gives a strong correlation, it can be inferred that there exists an anticorrelation between \hat{f}_P and T , because of the dependence of \hat{f}_{lim} on T in equation (2).

Similarly, instead of defining a variable threshold for \hat{f}_P we may define one for the duration T_x . It can be seen that if we define (see LP for more details)

$$T_{x,lim}(\hat{f}_P) = \frac{\eta\Delta t}{\hat{f}_P/\bar{f}_{lim} - 1} \quad (8)$$

then we have a two dimensional data set T_x and $T_{x,lim}$ with the truncation $T_x \geq T_{x,lim}$. Because of the dependence of $T_{x,lim}$ on \hat{f}_P , the test of correlation between T_x and $T_{x,lim}$ will amount to a test of anticorrelation between T_x and \hat{f}_P .

4.2.1. Peak Flux vs. Duration

Table 3 shows the results of several further tests utilizing \hat{f}_P and \hat{f}_{lim} or T_x and $T_{x,lim}$. The results of these tests are generally consistent with the \hat{f}_P – T tests. The only seemingly significant correlations between measures involving peak flux appear in the 64 ms column. If taken at face value, the fairly large positive values of t_1 for \hat{f}_P versus \hat{f}_{lim} would imply a significant dilation effect. However, this conclusion is suspect because t_1 for \bar{f}_P versus \bar{f}_{lim} gives a very significant correlation. This result is puzzling, because as stated above there should not be any correlation between the limiting flux set prior to the occurrence of a burst and the subsequent peak flux of the triggered burst. This correlation is unlikely to be caused by a few outliers, because it persists even when we divide the bursts up into subsets, either chronologically or in a random fashion.

There is the possibility that for some reason the transformation from counts \bar{C}_P to flux \bar{f}_P is biased in some systematic way. The relationship between the two quantities is

$$\bar{f}_P = \bar{C}_P / A_{eff}(\theta, \phi), \quad (9)$$

where $A_{eff}(\theta, \phi)$ is the effective area of the detector for the direction θ, ϕ of the burst, which depends mildly on its spectrum. In effect, this would be saying that the effective detector observing area or burst spectrum varies systematically with peak count rate or limiting count rate to produce the discrepant correlation. To test this hypothesis we show in Figure 5 plots of $A_{eff}(\theta, \phi)$ versus average peak count rate, with the bursts divided up by their limiting count rate. Due to the way the detector software operates, most of the bursts are triggered at one of three discrete values in \bar{C}_{lim} . It can be seen that only bursts with $C_{min} = \bar{C}_{lim}\Delta t = 60$ counts display a highly significant correlation. Currently we have no explanation for this result other than the generally unsatisfying explanation that it could be a statistical fluctuation. This behavior is not evidenced in $\Delta t = 1024$ ms data, and only very marginal evidence for it is present for $\Delta t = 256$ ms data. Table 4 summarizes the results.

It should be noted that the first five values in the last column of Table 3 give essentially the same values of t_1 . This result suggests that if the likely spurious effect (which could be due to an unfortunate and improbable fluctuation) were eliminated, then there will be little significant correlation remaining between \hat{f}_P and any of the durations. Using the other trigger durations as a guide, it can be seen that when the anomalous correlation is absent the t_1 values show no strong evidence for correlation.

4.2.2. Fluence vs. Duration

We now consider the correlation between fluence \mathcal{F} and duration T . As described in §3.2, from the BATSE data we can obtain \mathcal{F} , $\mathcal{F}_{lim} = \mathcal{F}C_{min}/C_{max}$, and some measure of duration, with $\mathcal{F} \geq \mathcal{F}_{lim}$. However, the truncation in the \mathcal{F} – T plane cannot be obtained directly from the data, so that we cannot directly test the correlation between these quantities. However, we do know that the burst selection process indicates that the threshold on the fluence will depend on the duration, with the exact relation depending on the pulse shape. Clearly for simple pulses $\mathcal{F}_{lim} \propto \hat{f}_{lim}T \propto \bar{f}_{lim}(\Delta t + T)$. Therefore, we can test the correlation between \mathcal{F} and T indirectly by considering the correlation between \mathcal{F} and \mathcal{F}_{lim} . The final row in Table 3 shows the result of the test between fluence \mathcal{F} and \mathcal{F}_{lim} , which indicated the presence of a significant correlation. Since the fluence limit becomes approximately proportional to T for long duration events and is not correlated with T for short duration events, a positive correlation between fluence and fluence limit would indicate a positive correlation between fluence and duration, especially for long duration bursts.

Taken at face value, the large positive test values indicate a highly significant correlation, which is in the opposite sense of that expected from the cosmological time dilation. It should be cautioned that the interpretation of these fluence results involves many difficulties. For example, fluence measures are much more sensitive to background subtraction than any kind of peak flux measure; it would not be hard to imagine further systematic biases that might affect the test results. Furthermore, changes in the burst spectrum with time could have a significant effect on the fluence while not mattering much for the determination of the peak flux. The fluence and the peak flux are also affected slightly differently by redshifting due to the extra time factor in the fluence, although this difference would not be enough to account for the positive results we find.

Assuming these issues can be ruled out as the cause for the positive test results, inspection of Figure 4 shows that even the lowest value of $t_1 = +2.28$ is inconsistent with any time dilation (even the lowest factor of 1.3) and implies a strong correlation between \mathcal{F} and T . Our finding that the slope of the $\log N$ – $\log \mathcal{F}$ has as a break in it that is as sharp or sharper than the break in the $\log N$ – $\log \hat{f}_P$ distribution (PL) would seem to indicate that the distribution of total radiant energy could be narrower than the distribution of peak luminosity. On this basis alone the fluence would seem to be the better “standard candle,” but it does not show the anticorrelation between fluence and duration expected for a static population of standard candles. One simple possibility is that bursts might not be of cosmological origin. However, this result does not rule out cosmological scenarios because any relation between \mathcal{F} and T and the $\log N$ – $\log \mathcal{F}$ curve can be fitted by invoking

an appropriate evolution of distributions, luminosities, or number density of bursts. In fact, for a cosmological population of bursts to be consistent with both the $\log N$ - $\log \mathcal{F}$ and the time dilation test results evolution appears to be required.

5. Conclusion

We have searched for time dilation effects in the BATSE 3B data by defining several measures of burst strength and duration and performing a nonparametric correlation test. Our treatment differs from previous treatments in that our test can account for nontrivial data truncations due to observational selection biases, allowing us to use a larger sample of bursts. Table 5 shows the percentage of all 1122 bursts in the BATSE 3B catalog that were able to be used in each test. These percentages should be compared to the approximately 20% used in previous time dilation tests.

The conclusions of this paper are:

1. We have confirmed that the observational definition of duration can have a major influence on correlation test results, and suggest effective duration (fluence divided by peak energy flux) as an appropriate measure for use in time dilation tests.
2. A nonparametric rank statistic test (Efron & Petrosian 1992) was used to overcome data truncation effects resulting from a short duration bias in the BATSE data. These tests utilize a greater number of bursts (up to 46% of the 1122 triggered bursts) than previous investigations and extend the test to short durations.
3. Test results for the correlation between duration and peak flux are consistent with no correlation, but are not currently sensitive enough to rule out the expected weak correlations. If existing trends continue, the 4B catalog may contain enough data to rule out a factor of 2 dilation. Ruling out a factor of 1.3 dilation would probably require a factor of six more bursts than currently available.
4. There appears to be slight evidence for different correlation properties for short and long duration bursts, but this evidence is not statistically compelling.
5. An indirect test of the correlation between fluence and duration indicates a positive correlation, which is inconsistent with simple no-evolution cosmological scenarios if fluence is a better standard candle than the peak flux. This point is discussed more completely in PL.

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	\bar{f}_P versus:		
	T_{90}	T_{50}	$T_{eff}(T_{50})$
64 ms (296 bursts)	1.02	-0.282	-0.406
256 ms (314 bursts)	2.01	1.04	0.977
1024 ms (382 bursts)	2.61	1.54	0.853
	$\hat{f}_P(T_x)$ versus:		
	T_{90}	T_{50}	$T_{eff}(T_{50})$
64 ms (296 bursts)	0.834	-0.444	-0.565
256 ms (314 bursts)	1.77	0.748	0.666
1024 ms (382 bursts)	2.02	0.829	0.356

Table 1: Values of t_1 for the correlation between peak fluxes and durations for all available BATSE bursts.

	\bar{f}_P versus:					
	T_{90}	T_{50}	$T_{eff}(T_{50})$	\hat{T}_{90}	\hat{T}_{50}	$T_{eff}(\hat{T}_{50})$
64 ms (181 bursts)	0.134	-1.20	-1.56	-1.22	-1.20	-1.56
256 ms (242 bursts)	1.09	0.153	0.113	-0.728	-0.266	0.110
1024 ms (288 bursts)	1.42	0.325	-0.142	-0.708	-0.334	-0.227
	$\hat{f}_P(T_x)$ versus:					
	T_{90}	T_{50}	$T_{eff}(T_{50})$	\hat{T}_{90}	\hat{T}_{50}	$T_{eff}(\hat{T}_{50})$
64 ms (181 bursts)	0.102	-1.23	-1.61	-1.28	-1.28	-1.60
256 ms (242 bursts)	1.08	0.113	-0.031	-0.703	-0.316	-0.033
1024 ms (288 bursts)	1.21	0.010	-0.287	-0.855	-0.582	-0.342

Table 2: Values of t_1 for the correlation between peak fluxes and durations for subsamples of bursts with durations measured by Norris.

	1024 ms (514 bursts)	256 ms (403 bursts)	64 ms (417 bursts)
\bar{f}_P vs \bar{f}_{lim}	0.812	1.74	3.01
$\hat{f}_P(T_{50})$ vs $\hat{f}_{lim}(T_{50})$	-0.434	1.43	2.91
$\hat{f}_P(T_{eff}(T_{50}))$ vs $\hat{f}_{lim}(T_{eff}(T_{50}))$	-0.266	1.60	3.11
T_{50} vs $T_{50,lim}(\bar{f}_P)$	-1.21	1.03	0.879
$T_{eff}(T_{50})$ vs $T_{eff,lim}(\hat{f}_P)$	-1.01	0.953	0.895
\mathcal{F} vs \mathcal{F}_{lim}	4.26	3.35	2.28

Table 3: Values of t_1 for various pairs of parameters indirectly testing time dilation.

Δt	C_{min}	t_1	number in sample
64 ms	60	3.87	39
64 ms	66	1.38	275
64 ms	71	0.737	81
256 ms	121	2.24	45
256 ms	132	-0.243	259
256 ms	143	1.44	84
1024 ms	242	0.513	88
1024 ms	264	0.512	299
1024 ms	286	-0.266	99

Table 4: Tests of correlation between $A_{eff}(\theta, \phi)$ and \bar{C}_P , for subsets with different values of $C_{min} = \bar{C}_{lim}\Delta t$.

	1024 ms	256 ms	64 ms
\hat{f}_P or \bar{f}_P vs T_x	34	28	26
\hat{f}_P vs \hat{f}_{lim}	46	36	37
T_x vs T_{lim}	46	36	37
\mathcal{F} vs \mathcal{F}_{lim}	46	36	37

Table 5: Percentages of all 1122 bursts available for each test.

Fig. 1.— Comparison of the various definitions of duration. *Top panel:* The ratio of T_{90} to $T_{eff}(T_{50})$ as a function of \bar{C}_P/\bar{C}_{lim} . *Center panel:* The ratio of T_{90} to T_{50} as a function of \bar{C}_P/\bar{C}_{lim} . The dotted line marks the value of the ratio 9/5. *Bottom panel:* The ratio of T_{50} to $T_{eff}(T_{50})$ as a function of \bar{C}_P/\bar{C}_{lim} . In all three panels, the solid lines indicate the best power law fits through the points and are meant to give an indication of the trends. The $\Delta t = 64$ ms trigger timescale was used for all of these.

Fig. 2.— Correlation test results for several measures of peak flux versus duration. The left panel shows the t_1 values for the correlation between \bar{f}_P and T_{90} (solid line), T_{50} (dotted line), and $T_{eff}(T_{50})$ (dashed line) for three different trigger timescales; t_1 is a test statistic which measures the deviation of the data from the null hypothesis that the variables are uncorrelated (see §2 for details). The right panel is the same except with $\hat{f}_P(T_x)$ instead of \bar{f}_P and gives systematically lower values of t_1 , showing the effect of the short duration bias. The larger values of t_1 for T_{90} in both panels show the effect of the signal-to-noise bias.

Fig. 3.— *Top panel:* Scatter plot of $T_{eff}(T_{50})$ versus $\hat{f}_P(T_{eff})$. The line shows the truncation boundary due to the duration bias. *Center panel:* The t_1 statistic as a function of minimum $T_{eff}(T_{50})$, where t_1 is a measure of the likelihood of a significant correlation. The leftmost value of t_1 includes all the bursts, while all subsequent values include only bursts down to some minimum $T_{eff}(T_{50})$. The dotted line shows a function proportional to $\sqrt{m/M}$, where m is the number of bursts in the subsample and $M = 296$ is the total number of bursts. *Bottom panel:* Same as in the center panel, except that t_1 is plotted as a function of maximum $T_{eff}(T_{50})$, so that the rightmost value includes all of the bursts.

Fig. 4.— Distributions of t_1 , a statistical measure of the probability that the data are correlated, obtained from simulations. The solid histogram shows the reference simulation with no correlation. The long dashed, dotted, and short dashed histograms show the t_1 distributions for simulations of correlated variables corresponding to time dilation factors of 1.3, 1.7, and 2.0. The vertical lines mark the average value of t_1 for each set of 1000 simulations. The distributions have been shifted vertically for clarity.

Fig. 5.— Scatter plots of the effective observing area versus average peak count rate, separated by limiting count rate, for the $\Delta t = 64$ ms sample. From top to bottom, the plots are for $\bar{C}_{lim}\Delta t = 60, 66$, and 71 , respectively. The lines are the best straight line fits of $A_{eff}(\theta, \phi) = m \log \bar{C}_P + b$.









